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Coal consumption and industrial production nexus in USA: Cointegration with two unknown structural breaks and causality approaches

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ABSTRACT

The paper investigates the causality relationships among industrial production index, coal consumption and employment in industrial sector for the period of 1973:1–2011:10 in USA. After noticing that there are breaks in the regression model, the Hatemi-J test for cointegration is employed to the cases that take into account two possible regime shifts. It is concluded that there is a long run relationship between industrial production and industrial coal consumption with the breaks at 1983:4 and 1998:4. We found a negative relationship between coal consumption and industrial production for the period of 1973:1–1983:4 and positive relationship for 1983:5–1998:4 period. For the last period that covers 1983:5–2011:10, the cointegration relationship turned to negative. In addition, the results show that causal relationship between coal consumption and industrial production changes over time.

Contents

1. Introduction

Over the past three decades the relationship between energy consumption and economic growth has been a major issue of debate among economists and policy makers. Although coal is the major element for the industrial revolution in the world, the environmental consequences of the sustained use of coal has drawn into question of long-term viability of coal in light of the emergence of cleaner and alternative energy sources [1]. Countries that benefited from their coal reserves in the 19th century are now industrialized countries. Coal keeps its major role because it has high density, low cost and ease of combustion but its use produces several types of emissions that adversely affect the environment. Coal consumption accounted for 37% of the total US emissions of carbon dioxide released into the Earth's atmosphere in 2010 [2].

The entry into force of the Kyoto Protocol to the United Nations Framework Convention on Climate Change has focused the political divide between the coalition of industrialized countries that support the Kyoto treaty and design to implement rigid climate policies, and the few industrialized countries that are unwilling to do so [3]. The Kyoto Protocol requires participating countries to reduce their carbon-dioxide emissions collectively to an annual average of about five percent below their 1990 level over the 2008–2012 period. Coal consumption patterns, especially in USA, will certainly be affected since the United States is home to the largest recoverable reserves of coal in the world. Will coal consumption in reduction cause economic shocks, if there is a causal relationship between coal consumption and economic

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growth? The causal relationship between economic growth and energy consumption has been studied in a large number of empirical studies. However, the results of these studies are mixed due to methodological differences and the time period chosen (see Ozturk [4]).

It is widely believed that discovery of peat lead to industrial revolution. But damages of fossil energy kinds on environment are not denied. So human being search alternative energy researches instead of fossil energies. But there are needs for benefit-cost analysis and this reveals some questions. That is both benefits and cost of coal consumption and alternative energy kinds should be well documented. In addition, coal is the largest source of electricity in the world. However, as renewable energy consumption increase in total consumption, they will tend to be more common than fossil fuels for electricity generation because of reducing CO2 emissions into the atmosphere and combating global climate change. Therefore, this paper examines coal consumption and GDP linkage.

If one looks at the studies between coal consumption and economic growth linkage, this question is answered by four different hypotheses: ***The growth hypothesis*** states to a situation in which coal consumption plays an important role in the economic growth process directly and/or as a complement to capital and labor. The growth hypothesis is supported, if uni-directional causality is found from coal consumption to economic growth. In this case, energy conservation policies aimed at decreasing coal consumption will have negative effects on economic growth. This view is also confirmed by Wolde-Rufael [5] for India and Japan. ***The conservation hypothesis*** signifies that economic growth is the dynamic which causes the consumption of energy sources. According to the conservation hypothesis there is a uni-directional causality running from economic growth to coal consumption. In this state, energy conservation policies which may prevent energy consumption will not have negative impact on economic growth which is confirmed by Jin-ke et al. [6] for China and Japan; Yang [7] for Taiwan; Jin-ke et al. [8] for Japan and China; Wolde-Rufael [5] for China and Korea. ***The feedback hypothesis*** implies a mutual relationship between coal consumption and economic growth. The feedback hypothesis is supported if there is bi-directional causality between coal consumption and economic growth. This hypothesis is checked by Yoo [9] for Korea; Li and Leung [10] for China Coastal and Central regions; Wolde-Rafael [5] for South Africa and USA; Apergis and Payne [1,11] for 20 OECD and 15 emerging countries. ***The neutrality hypothesis*** indicates that energy consumption does not affect economic growth. The absence of causality between energy consumption and economic growth provides evidence for the validity of the neutrality hypothesis. In this case, energy conservation policies devoted to reducing energy consumption will not impact economic growth which is confirmed by Jin-ke et al. [6] for India, South Africa and South Korea; Jin-ke et al. [8] for India and South Africa.

Although all of these studies contribute to investigating the relationship between coal consumption and industrial production, these analyses were introduced based on the assumption that the cointegrating vector remained the same during the period of study. However, there are many reasons to expect that the long-run relationship between the underlying variables might change such as shifts in the cointegrating vector can occur as a result of policy and regime changes and organizational or institutional evolution [12]. This can be an important issue if there are structural shifts in one or more of the energy consumption and industrial production series, since the existence of a structural break may disguise the true nature of any potential relationships among energy consumption, industrial production, capital and labor. The current paper builds on Hatemi-J [12] test for cointegration to the cases that take into account two possible regime shifts. In this test, the timing of each shift is unknown a priori and it is determined endogenously. The distributions of the tests are non-standard and generating new critical values via simulation methods. The size and power properties of these test statistics are estimated through Monte Carlo simulations, which demonstrate the tests have small size distortions and very good power properties.

The paper examines the causality relationships among industrial production index, coal consumption and employment in industrial sector for the period of 1973:1–2011:10 in USA. The paper is organized as follows: in Section 2, we describe model and data, in Section 3, we examine the links between coal consumption and industrial production, and give possible explanations for the econometric results that this research provides. We present the conclusions of our study and discuss policy implications in Section 4.

2. Model and data

In order to take into account the effect of two structural breaks on the parameters, we estimate the following regression model:

$$\begin{aligned}\ln \mathrm{IPI} = {} & \alpha_0+\alpha_1 D_{1t}+\alpha_2 D_{2t}+\beta_0 \ln \mathrm{CC}_t+\beta_1 D_{1t}\ln \mathrm{CC}_t+\beta_2 D_{2t}\ln \mathrm{CC}_t \\ & +\varphi_0 \ln \mathrm{L}_t+\varphi_1 D_{1t}\ln \mathrm{L}_t+\varphi_2 D_{2t}\ln \mathrm{L}_t+e_t \end{aligned} \quad (1)$$

where IPI, CC and L are industrial production index, coal consumption by industrial sector and employment in industrial sector, respectively. All variables were indexed as base year 2005 and monthly data used. Coal consumption by industrial sector measured as thousand tons and obtained from Monthly Energy Review which was provided by US Energy Information Administration. Industrial coal consumption was seasonally adjusted using X12. Seasonally adjusted industrial production index and seasonally adjusted industrial employment are taken from OECD database. A monthly data set is used for the period of 1973:1–2011:10 for USA.

D_{1t} and D_{2t} are binary variables defined as

$$D_{1t}=\begin{cases}0 \text{ if } t\le T_1\\ 1 \text{ if } t>T_1\end{cases} \text{ and } D_{2t}=\begin{cases}0 \text{ if } t\le T_2\\ 1 \text{ if } t>T_2\end{cases}$$

where T is sample size. T_1 denotes the period before the first break and T_2 denotes the period before the second break.

3. Methodology and results

In this study, the method which was used to estimate cointegration test with two breaks consists of three steps. First step is to test the unit roots. Finding breaks in the model is the second step. Finally the cointegration test is carried out.

Before testing for cointegration, all of the variables in the model should meet the condition of $I(1)$. For this purpose, Kwiatkowski et al. [KPSS, [13]] and Lee and Strazicich [14] unit root tests with two endogenous structural breaks were used. The findings indicated that each variable is integrated to the first order.

In the second step, breaks in the multiple linear regression models were analyzed as suggested by Bai and Perron [15,16] who concentrate on the multiple linear regression system

$$Y = X\beta + \overline{Z}\delta + u \quad (2)$$

where $Y=(y_1,\ldots,y_T)'$, $X=(x_1,\ldots,x_T)'$, $U=(u_1,\ldots,u_T)'$, $\delta=(\delta_1',\delta_2',\ldots\delta_{m+1}')'$ and $\overline{Z}$ is the matrix that diagonally partitions Z at $(T_1,\ldots,T_m)$. Estimation method of the model relies on the least squares principle. To detect breaks in regression model, Bai and

Perron [15,16] suggest supF, UDmax, WDmax and subF$_T$ ($\ell+1|\ell$) tests.

SubF test has null hypothesis of no structural break against alternative hypothesis. SubF test is defined as [15,16]

$$F_T(\lambda_1,\ldots,\lambda_k;q)=\frac{1}{T}\left(\frac{T-(k-1)q-p}{kq}\right)\hat{\delta}' R'^{(R\hat{V}(\hat{\delta})R')^{-1}}R\hat{\delta} \tag{3}$$

R is the conventional matrix that $(R\delta)'=(\delta_1'-\delta_2',\ldots,\delta_k'-\delta_{k+1}')$. $\hat{V}(\hat{\delta})$ is an estimate of the covariance matrix of $\hat{\delta}$ robust to serial correlation and heteroscedasticity.

UDmax and WDmax tests have the null hypothesis of no structural break versus an unknown number of breaks given some upper bound M, whereas all weight equal to unity for UDmax, q and the significance level of test determines weights of the WDmax. SubF$_T$ ($\ell+1|\ell$) tests the null hypothesis of ℓ breaks against the alternative that an additional break exists.

Table 1 indicates the results of break tests suggested by Bai and Perron [15,16]. Since all variables was found $I(1)$, first differences of the variables were used. Regression model consists of industrial production and coal consumption as independent variable.

SubF, UDmax and WDmax tests indicate that there are breaks in the multiple regression models, and sequential procedure chooses two breaks. So, according to Bai and Perron tests results, one can accept that there are two breaks in the model.

Since there are breaks in the regression model, cointegration test should include breaks. Third stage of analyses consists of cointegration test with two unknown structural breaks suggested by Hatemi-J [12]. To test null hypothesis of no cointegration, author uses three statistics: Z_α and Z_t tests developed by Phillips [17] and augmented Dickey–Fuller (ADF) test. The test statistics are based on the calculation of the bias-corrected first-order serial correlation coefficient estimate ($\hat{p}^\bullet$) Hatemi-J [12] defines $\hat{p}^\bullet$ as

$$\hat{p}^\bullet=\frac{\sum_{t=1}^{T-1}(\hat{e}_t\hat{e}_{t+1}-\sum_{j=1}^{B}w(j/B)\hat{\gamma}(j))}{\sum_{t=1}^{T-1}\hat{e}_t^2} \tag{4}$$

where $\hat{e}_t$ is the estimated value of at time t for the estimated model with T observations. $w(\cdot)$ is a function providing kernel weights meeting the standard conditions for spectral density estimators. B is the bandwidth number satisfying the conditions $B\to\infty$ and $B/T^5=O(1)$. $\hat{\gamma}(j)$ is an autocovariance function. The autocovariance function is defined as

$$\hat{\gamma}(j)=\frac{1}{T}\sum_{t=j+1}^{T}(\hat{e}_{t-j}-\hat{p}\hat{e}_{t-j-1})(\hat{e}_t-\hat{p}\hat{e}_{t-1}) \tag{5}$$

here $\hat{p}$ indicates the estimated values of the effect of $\hat{e}_{t-1}$ on . The Z_α test statistic is defines as

$$Z_\alpha=T(\hat{p}^\bullet-1) \tag{6}$$

The Z_t test statistic is defined as follows:

$$\frac{Z_t=(\hat{p}^\bullet-1)}{(\hat{\gamma}(0)+2\sum_{j=1}^{B}w(j/B)\hat{\gamma}(j))/\sum_1^{T-1}\hat{e}_t^2} \tag{7}$$

where $(\hat{\gamma}(0)+2\sum_{j=1}^{B}w(j/B)\hat{\gamma}(j)$ denotes the long-run variance of the residuals of a regression of $\hat{e}_t$ on $\hat{e}_{t-1}$ The distribution of these tests does not follow standard asymptotic distributions and using Monte-Carlo simulations new critical values are generated by Hatemi-J [12]. The test statistics are the smallest values of these three tests across all values for τ_1 and τ_2, with $\tau_1\in T_1=(0.15, 0.70)$ and $\tau_2\in T_2=(0.15+\tau_1, 0.85)$. These test statistics are as follows:

$$ADF^*=\inf_{(\tau_1/\tau_2)\in T}ADF(\tau_1,\tau_2) \tag{8}$$

$$Z_t^*=\inf_{(\tau_1/\tau_2)\in T}Z_t(\tau_1,\tau_2) \tag{9}$$

$$Z_\alpha^*=\inf_{(\tau_1/\tau_2)\in T}Z_\alpha(\tau_1,\tau_2) \tag{10}$$

Table 2 represents the results of Modified ADF, Z_t and Z_α cointegration tests.

According to results of Hatemi-J cointegration tests, one can conclude that there is a long run relationship between industrial production and industrial coal consumption. All of the tests reject the null hypothesis of no cointegration at the 1% significance level. The first structural break is found at 1983:4 and the second one is at 1998:4. The potential reason of the breaks may be the cost of coal consumption. Graph 1 illustrates the cost of coal receipts for the period of 1973–2010.

According to Graph 1, the cost of coal increases to 1983 and decreases to 1998 and again increases after 1998. Table 3 represents the estimated values of the parameters by Hatemi-J cointegration test.

All of the coal consumption parameter values are statistically significant. $\beta_0$ln CC covers the period of 1973:1–1983:4. In this period there is a negative relationship between coal consumption and industrial production. Since cost of coal increases in the same period, alternative energy sources may be used instead of coal in industrial sector. So while industrial production was increasing, coal consumption of industrial sector was decrease. $\beta_1$ln CC covers the period of 1983:5–1998:4; while cost of coal was decreasing, the relationship between coal consumption and industrial production was positive. $\beta_2$ln CC covers the period of 1983:5–2011:10. As cost of coal was increasing again in this period, the cointegration relationship between coal consumption and industrial production turned to negative. In short, cost of coal may affect the relationship between coal consumption and industrial production by changing decision to use of energy kind in industrial sector.

If two or more time-series are cointegrated, there must be Granger causality between them—either one-way or in both directions. However, the converse is not true. For the purpose of cross-check on the validity of our results, we used Toda and Yamamoto [18] and Hacker and Hatemi-J [19] procedures to test for Granger causality. The causality tests are applied as following steps:

1. Since two breaks found in the model, the data set divided into three sub periods. The first one covers the period of 1973M1–1983M4. The second period is between 1983M5 and 1998M4.

Table 1
Breaks in the multiple linear regression model.

SubF tests	Test stat	SubFT($\ell+1\|\ell$) test	Test stat
0 vs. 1	22.4294***	2 vs. 1	21.9776***
0 vs. 2	46396.1873***	3 vs. 2	6.5832
0 vs. 3	5770.3256***	4 vs. 3	5.4956
0 vs. 4	147821.2336***	5 vs. 4	0.0576
0 vs. 5	556066.1955***	Number of breaks	2
UDmax test	556066.1955***	WDmax test	1220962.4893***

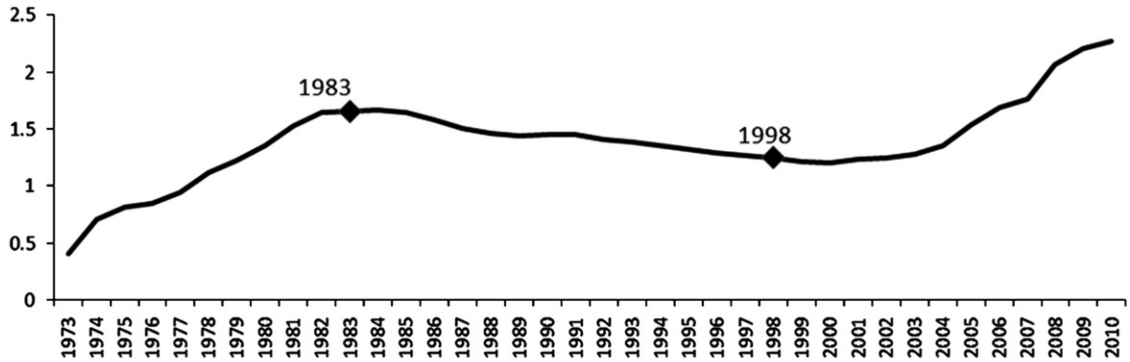
*** Denotes that the tests are significant at 1% level.

Table 2
Results of modified ADF, Z_t and Z_α cointegration tests.

	Test statistics	1% CV	%5 CV	10% CV
Modified ADF	−7.315***	−6.928	−6.458	−6.224
Z_t	−7.616***	−6.928	−6.458	−6.224
Z_α	−101.399***	−99.458	−83.644	−76.806

Note: critical values (CV) were taken from Hatemi-J (2008).

*** Denotes that the tests are significant at 1% level.



Graph 1. Cost of coal receipts.
Source: US IEA

Table 3
Estimated values of the parameters.

	Coefficient	**SE**	**t**
α_0	−0.618	0.552	−1.119
α_1	5.912	1.138	5.195
α_2	−0.274	1.025	−0.268
$\beta_0$ln CC	−0.247	0.040	−6.115
$\beta_1$ln CC	1.188	0.115	10.334
$\beta_2$ln CC	−1.450	0.108	−13.401
$\varphi_D$ln L	0.313	0.269	1.166
$\varphi_1$ln L	1.987	0.136	14.559
$\varphi_2$ln L	−1.886	0.268	−7.035

Note: the Newey–West heteroscedasticity and autocorrelation consistent standard errors are used.

Table 4
Causality test results.

Period	The Null hypothesis	Toda and Yamamoto test		Bootstrap critical values		
		MWALD	Prob.	%1	%5	%10
1973M1–	CC ≠ IPI	4.289807	0.1171	10.118	6.325	4.794
1983M4	IPI ≠ CC	3.575527	0.1673	10.128	6.353	4.806
1983M5––	CC ≠ IPI	5.045888	0.4103	15.968	11.583	9.603
1998M4	IPI ≠ CC	12.80620	0.0253	15.863	11.540	9.552
1998M5–	CC ≠ IPI	17.56180	0.0074	18.074	13.106	10.947
2011M10	IPI ≠ CC	36.99160	0.0000	18.857	13.629	11.361

Note: CC: coal consumption, IPI: industrial production index.

Time span of the last period is 1998M5–2011M10. It was tested for each of the time-series to determine their order of integration. KPSS test results indicated that integration order of most of the series is one ($I(1)$).

2. It was determined the appropriate maximum lag length for the variables in the VAR using SIC. The problem of serial correlation in the residuals was taken into account using LM test and lag length was increased until any autocorrelation issues was resolved. The appropriate maximum lag length for models which were divided into three periods is 2, 5 and 6, respectively. In addition, inverse roots of AR characteristics were investigated. The estimated models are dynamically stable except for the model which covers the first period.
3. VAR($p+d$) model set up as suggested by Toda and Yamamoto [18] and it is as follows:

$$IPI_t = \eta_0 + \sum_{i=1}^{k}\gamma_{1i}IPI_{t-i} + \sum_{j=k+1}^{d_{\max}}\gamma_{2j}IPI_{t-j} + \sum_{i=1}^{k}\varphi_{1i}\mathrm{CC}_{t-i} + \sum_{j=k+1}^{d_{\max}}\varphi_{2j}\mathrm{CC}_{t-j} + \sum_{i=1}^{k}\delta_{1i}\mathrm{L}_{t-i} + \sum_{j=k+1}^{d_{\max}}\delta_{2j}\mathrm{L}_{t-j} + \varepsilon_{1t} \quad (11)$$

$$\mathrm{CC}_t = \tilde{\eta}_0 + \sum_{i=1}^{k}\tilde{\gamma}_{1i}\mathrm{IPI}_{t-i} + \sum_{j=k+1}^{d_{\max}}\tilde{\gamma}_{2j}\mathrm{IPI}_{t-j} + \sum_{i=1}^{k}\tilde{\varphi}_{1i}\mathrm{CC}_{t-i} + \sum_{j=k+1}^{d_{\max}}\tilde{\varphi}_{2j}\mathrm{CC}_{t-j} + \sum_{i=1}^{k}\tilde{\delta}_{1i}\mathrm{L}_{t-i} + \sum_{j=k+1}^{d_{\max}}\tilde{\delta}_{2j}\mathrm{L}_{t-j} + \varepsilon_{2t} \quad (12)$$

4. Granger non-causality test was applied using modified WALD test.
5. Normality of residual was tested using the Doornik–Hansen multivariate normality test. According to normality test results, the residuals are not normally distributed. However, if the residual is not normally distributed, the MWALD test based on the asymptotic critical values does not have correct size properties as shown by Hacker and Hatemi-J [19]. The authors develop a test method based on leveraged bootstrap simulation techniques that produces more precise critical values. Furthermore the leveraged bootstrap causality test ensures that the presence of heteroscedasticity does not affect the accuracy of estimated results (For details see Hacker and Hatemi-J [19]).
6. Finally Hacker and Hatemi-J bootstrap causality test was run and its results compared to results of Toda and Yamamoto test. Table 4 illustrates the causality test results.

According to Toda and Yamamoto procedure and Hatemi-J test, the results are robust against non-normality of the residual. For the first period there is no causal relationship between coal consumption (CC) and industrial production index (*IPI*). This findings support the neutrality hypothesis. In the second period, there is a causal relationship from IPI to CC. So conservation hypothesis is supported in the second period. In the last period, there is bidirectional causal relationship between CC and IPI as suggested by the feedback hypothesis. All of these results indicate that causal relationship between coal consumption and industrial production changes over time. Thus, the breaks in this relationship should take into account in the empirical studies.

4. Conclusion

This paper focuses on the coal consumption and its effect on industrial production. There are many reasons to expect that the

long-run relationship between the underlying variables might change. To test parameter changes between industrial production and its coal consumption, the method suggested by Bai and Perron [16] was used. Since the sequential procedure choose two breaks in the parameters, the model is tested by using the Hatemi-J [12] cointegration test which takes into account two possible regime shifts. The size and power properties of these test statistics are estimated through Monte Carlo simulations which demonstrate the tests have small size distortions and very good power properties.

The paper investigates the causality relationships among industrial production index, coal consumption and employment in industrial sector for the period of 1973:1–2011:10 in USA. According to results of Hatemi-J cointegration tests, there is a long run relationship between industrial production and industrial coal consumption. The first structural break is found at 1983:4 and the second one is at 1998:4. The potential reason of the breaks may be the cost of coal consumption. While the cost of coal is increasing (decreasing) there is negative (positive) relationship between coal consumption of industrial sector and industrial production. The cost of coal may affect the relationship between coal consumption and industrial production by changing decision to substitution between energy kinds in industrial sector. So, price of coal and relative price of energy kinds may be one of the main determinants of the relationship between energy consumption and economic growth.

If two or more time-series are cointegrated, there must be Granger causality between them. For the purpose of cross-check on the validity of our results, we used Toda and Yamamoto [18] and Hacker and Hatemi-J [19] procedures to test for Granger causality. Findings indicate that causal relationship between coal consumption and industrial production changes over time. For the first period there is no causal relationship between CC and IPI. In second period there is a causal relationship from IPI to CC. In the last period, there is bidirectional causal relationship between CC and IPI as suggested by the feedback hypothesis. Thus, regime shifts should take into account in empirical studies and regulation in energy markets. Our results indicate that coal consumption should be reduced in the first two periods, since this policy does not affect the industrial production. But in the last period, conservation of coal consumption affects the industrial production. Therefore, conservation policy is not plausible option in the last period.

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